Marine Resource Economics, Volume 11, pp. 143–155 0738-1360/96 $3.00 + .00

Substitutability Among Species in the Japanese Tuna Market: A Cointegration Analysis

SHEKAR BOSE
ALISTAIR McILGORM
Australian Maritime College

Abstract *This paper examines two sets of hypotheses concerning the existence and the cause of the long-run inter-species price relationships in the Japanese tuna market. A shock variable is introduced into the system to determine the degree of influence on the price relationships as well as the magnitude of the power in explaining the variation in prices of tuna species. Although in most cases the coefficient estimates of the shock variable are statistically significant, overall, the variable does not have significant explanatory power in both bivariate and multivariate regressions. We also find that the degree of substitutability between bigeye and albacore is substantially lower than the degree of substitutability between bigeye and yellowfin and, yellowfin and albacore.*

Key words Cointegration, macroeconomic shocks, price relationships, seasonality, substitutability.

Introduction

The study of market price behavior is important from a number of perspectives. One of the most important contributions of observed market prices is that they transmit information among market participants and the informational efficiency provided by the observed market prices is the prerequisite to achieve allocative efficiency. Informational inaccuracy regarding prices may distort marketing decisions and lead to inefficient inter-market commodity movements.

In the past few years, in the area of fisheries economics, a number of analytical and empirical studies have been devoted to the investigation of price determination (Kirman 1992; Pascoe, *et al.* 1987), price linkage and spread between varieties of domestic markets (Squires 1986; MacIntosh, *et al.* 1988), cross-country market integration (Squires, *et al.* 1988), market linkage between high- and low-valued fish species (Gordon, *et al.* 1993), and price interrelationship among different fish products (Hannesson 1994a, 1994b).

In this study we wish to follow a different approach to investigate inter-species price relationships and hence the degree of substitutability between species in the Japanese tuna market. The approach comprises two sets of hypotheses testing within the framework of time series analysis. First, we test the null hypothesis of the non-existence of long-run relationships against the alternative of existence of such rela-

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Shekar Bose is Lecturer, Faculty of Fisheries and the Marine Environment, Australian Maritime College, P.O. Box 986, Launceston, Tasmania, Australia 7250. Alistair McIlgorm is Associate Director (Fisheries), Faculty of Fisheries and the Marine Environment, Australian Maritime College.
The authors would like to thank the Associate Editor of this journal and two anonymous reviewers for useful and helpful comments and suggestions made on an earlier draft of this paper. The assistance of R. Bond of the Australian Maritime College with the presentation of this paper and the financial support of AMC Search Ltd. are gratefully acknowledged. A special thanks is extended to Mr. Tony Kingston, Senior Economist, South Pacific Forum Fisheries Agency for the provision of the price series data on the Japanese tuna market. The usual disclaimer applies.

tionships in both bivariate and multivariate contexts taking all variables into consideration. In the second stage, we also wish to test the null hypothesis of such long-run price relationships are due to substitutability between species against the alternative hypothesis that such relationships are due to the influence of other macroeconomic factors, provided that the null hypothesis is rejected in the first stage. The second set of hypotheses test whether co-movement of prices is a result of the spill-over effects of any macroeconomic shock to the economy (Pindyck and Rotemberg 1990).

Most of the market studies generally concentrate on understanding the behavior of the market participants. Substitutability between tuna fish species depends upon the consumer's perception of the relative importance of attribute categories (*e.g.* price, size, color, appearance, fat and moisture contents, flesh, texture, and species). For instance, the *sashimi* market generally demands high quality large tuna species, such as bluefin, bigeye, and to a lesser extent yellowfin, and thereby commands a high price. The consumers in the *sashimi* market are very sensitive to quality which is usually affected by product attributes (Ashenden and Kitson 1987). Thus, to be substitutable, two species would need to be perceived by the consumer as being of similar high quality as well as closely priced (Wessells and Anderson 1992).

Review of Related Studies

In their study on Rungis fish market in Paris, Gordon, *et al.* (1993) tried to identify whether there exists a formal market linkage between high-valued (salmon and turbot) and low-valued (cod) fish species. Using both the Engle-Granger (1987) and Johansen (1988) cointegration test procedures they found no evidence to justify that the market for salmon is linked to the markets for turbot or cod.

Using cointegration technique, Hannesson (1994a, 1994b) examined the issue of substitutability between different fish products (fresh and/or frozen) for the European community and the U.S. markets. Based on his findings, he argued that the existing cointegration between price series of different frozen fish products is due to the high degree of substitutability between those products. It should be mentioned that he used the import price data for the investigation.

Owen and Troedson (1994b) investigated the relationship between the prices of five tuna species at Yaizu market in Japan for the period 1976(1)–1992(8) using the vector autoregression technique. In their study they found all price variables are first difference stationary, that is, integrated of order one [I(1)]. As a next step, they employed Johansen's procedure to examine the cointegrating relationship between the price variables, but the test results did not show any evidence of cointegration at the conventional levels of significance.

Methodology

To serve the aforementioned purpose of our study we use the time series techniques of cointegration proposed by Engle and Granger (1987) and Johansen (1988).

Preliminary Tests: Stationarity (or Unit roots)

The stationarity of a time series variable simply implies that the series possesses the desirable linear statistical properties, such as, time invariant conditional mean, variance, and autocovariance (Granger 1986).

In the time series literature, non-stationarity behavior of the time series variable is an established fact. Therefore, prior to examining the long-run cointegrating relationships between variables, a test for the stationarity of the individual series should be performed to avoid spurious results (Granger and Newbold 1974).

It is also reasonable to expect that seasonally unadjusted price series may contain seasonal components. Thus, tests for regular non-seasonal unit roots should be preceded by the tests for seasonal unit roots. Combining both the factors of seasonality and non-seasonality, the definition of integration of univariate series becomes as follows: a price series p_t is said to be integrated of order (d, s), denoted as $p_t \sim I(d, s)$, if the series has a stationary, invertible, non-deterministic autoregressive moving average representation after one-period differencing 'd' times and seasonal differencing 's' times (Dolado, *et al.* 1990).

In testing for seasonal unit roots we follow the test procedures provided by Beaulieu and Miron (1993) for monthly data in which the null hypothesis of I(1, 1) is tested against the alternative hypothesis of I(1, 0) (for the outline of the test procedure see Beaulieu and Miron 1993). Critical values of tests statistics are given in table A.1 of Beaulieu and Miron (1993).

To test for non-seasonal unit roots, we consider the Augmented Dickey-Fuller (ADF) test procedure and run the following regressions:

$$\Delta X_t = \alpha + \gamma X_{t-1} + \sum_{j=1}^{k} \delta_j \, \Delta X_{t-j} + e_t \tag{1}$$

$$\Delta X_t = \alpha + \beta t + \gamma X_{t-1} + \sum_{j=1}^{k} \delta_j \, \Delta X_{t-j} + e_t \tag{2}$$

where, x_t represents the proposed variable, t stands for time, Δ is the difference operator, and e_t is the white noise residual. To detect the stationarity of the individual series one should decide between the hypotheses H_0: $\gamma = 0$ (non-stationary) and H_A: $\gamma \neq 0$ (stationary). The t-ratios obtained from the Ordinary Least Square (OLS) for the estimated coefficient 'γ' are then compared to the tabulated critical values given in Fuller (1976, p. 373). Sufficient lag terms of the dependent variables are included into the regressions to make the residuals white noise. In addition to the ADF test, we have also followed the F-test procedure suggested by Dickey and Fuller (1981).

Cointegration and Associated Tests

The logic behind the notion of cointegration is to establish statistically sound long-run relationships between economic time series. An important implication of cointegration between prices of two interrelated commodities is that pairs of such variables should not diverge from one another to a great extent in the long run (Granger 1986). The underlying reason for this is that economic forces (such as market forces) will play an important role to prohibit a persistent deviation from their relevant long run behavioral path. For instance, if price of a commodity is reasonably higher than the prices of other substitute commodities in the same market, then consumers will switch their consumption behavior to substitute commodities. As a result, the price of the first commodity will decline following the law of demand (Hannesson 1994a). This process will continue until the price gap between the two related commodities is reasonably narrowed down so that consumers will stop to switch their consumption behavior form the first commodity to another substitute commodity. Thus, we can see that prices of the substitute commodities in the same market should not diverge from each other to a great extent in the long-run. In other words, prices of substitute commodities in the same market should be cointegrated.

Table 1
Results of the Seasonal Unit Roots Tests

Frequency:	0	π	$\pi/2$		$2\pi/3$		$\pi/3$		$5\pi/6$		$\pi/6$		$\pi/2$	$2\pi/3$	$\pi/3$	$5\pi/6$	$\pi/6$		
Statistic:	π_1	π_2	π_3	π_4	π_5	π_6	π_7	π_8	π_9	π_{10}	π_{11}	π_{12}	$F_{3,4}$	$F_{5,6}$	$F_{7,8}$	$F_{9,10}$	$F_{11,12}$	LM(12)	$Q(12)$
P_A	3.32	–4.90	–3.90	–5.42	–4.40	2.56	–1.28	–7.02	–6.08	2.18	1.30	–6.48	24.67	13.65	26.80	21.14	21.19	0.99	2.72
P_B	4.14	–4.98	–2.03	–6.30	–3.15	4.68	–1.56	–6.91	–4.91	4.34	3.33	–6.77	22.81	15.93	26.65	20.63	26.29	0.34	1.67
P_Y	4.11	–3.82	–3.85	–6.28	–1.73	5.47	–2.40	–6.15	–5.83	2.48	0.74	–8.76	30.18	16.14	23.18	20.13	38.61	1.04	2.89

Note: P_A, P_B, and P_Y represent the prices of albacore, bigeye, and yellowfin respectively. The regressions include an intercept, eleven seasonal dummies, and a trend. Critical values of the test statistics are given in table A.1 of Beaulieu and Miron (1993, pp. 325–26). The last two columns represents the values of Lagrange Multiplier (LM) and Box-Pierce-Ljung 'Q' (with lag in the parentheses) tests for residual autocorrelation.

The simple statistical definition of cointegrated series is as follows. If, for a pair of time series variables P_{it} and P_{jt} ($i \neq j$), each of which is first difference stationary [*i.e.* I(1)], there exist a vector $[1, -\beta]^T$, such that the linear combination of the two series: $P_{it} - \alpha - \beta P_{jt} = Z_t$ is $I(0)$, then the series P_{it} and P_{jt} are said to be cointegrated of order (1, 1). The linear relationship between the series is called cointegrating regression with 'β' as a cointegrating parameter. The term Z_t is called an equilibrium error and it measures the extent of short run deviation from the long run equilibrium path. This idea could also be extended to a multivariate context. Thus, testing for cointegration involves an examination of the residuals from the cointegrating regression to satisfy the requirement that the residual from the regression is I(0). We use the same ADF test procedure outlined above to examine the order of integration of the residual from our bivariate and multivariate cointegrating regressions. Although the Engle-Granger procedure is used in determining the cointegrating relationship between our proposed variables is computationally simple, it suffers from some problems, such as, arbitrary selection of dependent variables and failure to identify the number of cointegrating vectors for multivariate case. To avoid these problems we employ the recently developed likelihood ratio (LR) test procedure by Johansen (1988) and Johansen and Juselius (1990). Their proposed LR test for the hypothesis that there are at most 'r' cointegrating vectors is given by:

$$\mathrm{LR} = -T\sum_{i=r+1}^{N} \ln(1 - \hat{\lambda}_i) \tag{3}$$

where $\hat{\lambda}_{r+1} \ldots \hat{\lambda}_N$ are the $N - r$ smallest squared canonical correlation coefficients between the residuals obtained by first regressing ΔX_t (t = 1, 2, ..., T) on its lagged differences, and then regressing X_{t-k} on the same regressand. Here 'X' represents a vector of N variables of interest. The asymptotic distribution of the LR test statistic is given by a multivariate version of the Dickey-Fuller distribution (Johansen and Juselius 1990). Full details of theoretical backgrounds and application guide of the Johansen's test procedure are provided in Dickey and Rossana (1994).

Brief Discussion on the Japanese Tuna Market

The Japanese tuna market has two distinct segments: fresh and frozen. From an international perspective, the unique feature of the market is the dominance of *sashimi* which is the preferred market form in Japan. Next to the highest priced southern bluefin tuna (*Thunnus maccoyi*), yellowfin (*Thunnus albacares*) and bigeye (*Thunnus obesus*) are usually used for *sashimi* products in Japan (Williams 1986). However, albacore (*Thunnus alalunga*) is a white meat tuna and is commonly used for canned tuna production.

Of the major tuna fishing nations Japan had the world's highest catch of 776,600 mt (preliminary estimate) in 1993. Moreover, Japan is regarded as one of the world's major markets for tuna products. In 1993, Japan also imported 336,978 mt of fresh and frozen tuna from other countries of which approximately 52.21%, 27.41%, and 0.72% were yellowfin, bigeye, and albacore respectively (Peckham 1995).

Yamamoto (1994) and Owen and Troedson (1994a) mentioned recently that in Japan, bigeye and albacore have been appreciated as substitutes for bluefin due to the supply scarcity of bluefin in Japan.

The Japanese purse seine and longline tuna fleets land their products at the Yaizu port (which is one of the largest wholesale entry points in Japan) after voyages to the Western Pacific and other tuna grounds. However, it should be noted that in the Japanese system producers usually land their product into the competitive wholesale auction markets where the price is likely to be determined by the interaction between the demand and supply (Pascoe, *et al.* 1987).

Data Source

We employ the monthly average market prices (Yen/kg) data for three different tuna species—yellowfin, bigeye, and albacore—which covers the period 1975 (1) to 1994 (11). We obtained the Yaizu price series from the South Pacific Forum Fisheries Agency, Solomon Islands, which has been supplied by the Japanese authorities and is used as the basis for negotiating contractual fishing arrangements in the waters of South Pacific nations.[1] It should be mentioned here that the species prices are for tuna captured by the longline fishing method as opposed to alternatives such as purse seine and pole and line fishing. Furthermore, the prices are for frozen product. Data on consumer price index (1990 = 100) and the industrial production index (*IP*) of finished consumer non-durable goods (1985 = 100) are taken from OECD Main Economic Indicators. Data on the index of industrial production are seasonally adjusted and are used as a proxy for macroeconomic shocks. Sufficient care has been taken to obtain a consistent data series. All variables are in real and logarithmic form.

Discussion of the Results

Table 1 presents the results of the seasonal unit root tests for price variables. Being seasonally adjusted, the series of industrial production index is excluded from the test. The regression equations include an intercept, time trend, and eleven seasonal dummy variables. As both Lagrange Multiplier (LM) and Box-Pierce-Ljung (denoted as Q) statistics fail to detect the presence of residual autocorrelation in the models, we do not include any lag of the dependent variables in our models. Based on our tests results, we reject the null hypothesis of seasonal unit roots for frequency π at the 1% level of significance, but we are unable to reject the null hypothesis in case of zero frequency, as all π_1 statistics are positive. For other frequencies (excluding $\pi/2$ to avoid the situation of non-existence of unit roots) and when k is even, we reject the null hypothesis at the 5% level of significance. We also employ the 'F' test suggested by Beaulieu and Miron (1993). The test results strongly reject the null hypothesis as all the calculated F-values are higher than the critical values along with the fact that π_2, and at least one member of each of the following subsets of test statistics $\{\pi_3, \pi_4\}$, $\{\pi_5, \pi_6\}$, $\{\pi_7, \pi_8\}$, $\{\pi_9, \pi_{10}\}$, and $\{\pi_{11}, \pi_{12}\}$, are significantly different from zero. Thus, the overall test results indicate that the price series do not contain any seasonal unit roots at any seasonal frequency other than zero.

The results of the non-seasonal unit root tests on all the variables are presented in table 2. Sufficient lag terms of the dependent variable have been added to the models (1) and (2) to whiten the residual. For the level form of all variables, test results do not allow us to reject the null hypothesis of non-stationarity at the 5% level of significance. But for the first difference case, the null hypothesis of non-stationarity is rejected at the 1% level in favor of the alternative of stationarity for all variables. Thus, based on our test results, we can say that all variables are stationary in their first differences. In other words, they are integrated of order one [*i.e.* I(1)].

Table 3 reports the results of the pair-wise cointegration regressions and the as-

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[1] Data are confidential and would not be available from the authors.

Table 2
Results of the Regular Unit Root Tests

	P_A	P_B	P_Y	IP
Level form				
ADF-test				
with constant	–0.87	–1.81	–2.37	–2.01
with constant and trend	–3.34	–2.20	–3.16	–1.98
F–test				
ϕ_1	1.10	2.05	2.99	2.34
ϕ_2	4.25	2.09	3.70	1.63
ϕ_3	5.61	2.71	5.37	2.16
First difference form				
ADF–test				
with constant	–10.27	–9.35	–9.82	–4.70
with constant and trend	–10.26	–9.33	–9.82	–4.73
F–test				
ϕ_1	52.71	43.72	46.27	11.08
ϕ_2	35.12	29.07	32.15	7.49
ϕ_3	52.67	43.50	48.22	11.23

Note: *IP* represents the index of industrial production. Critical values of ADF (Augmented Dickey-Fuller) statistic at 1% and 5% levels are –3.46 and –2.88 (for regression including constant), and –3.99 and –3.43 (for regression including both constant and trend) respectively, (Fuller 1976, p. 373). Critical values of ϕ_1, ϕ_2, and ϕ_3 (F-tests) at 1% and 5% levels are 6.52 and 4.63, 6.22, 4.75, and 8.43 and 6.34, respectively (Dickey and Fuller 1981, p. 1063).

sociated stationary tests of the residuals. The first set of regressions [*i.e.* (*a*), (*b*), and (*c*)] examines the cointegration relationships between price variables.[2] On the other hand, the second set of regressions [*i.e.* (*d*), (*e*), and (*f*)] examines the relationships between one of the price series and the shock variable IP. We consider two types of ADF tests. The first type includes only a constant and the second type includes both a constant and trend. For regression (*a*), first type of the ADF test results is significant at the 5% level, while the second type is significant at the 1% level. For regression (*b*), the ADF test results for both types are significant at the 1% level. Although the ADF statistic is not significant at the 5% level for the first type of regression (*c*), it becomes significant at the 5% level for the second type.

Considering the second set of regressions, the first type of ADF test statistics are insignificant for regressions (*d*) and (*e*), but they are significant at the 5% level for the second type of ADF tests. On the other hand, for regression (*f*) both types of ADF tests are significant at the 5% and 1% levels respectively. The Johansen's LR test results for both sets of regressions indicate the presence of one cointegrating vector at the conventional levels of significance. Thus, our statistical test results allow us to reject the null hypothesis of non-cointegration in favor of the alternative of cointegration. It is worth mentioning that for each regression, except (*d*), the independent variable is statistically significant at the conventional level. Furthermore, for the second set of regressions, R^2s (a measure of explanatory power of the regression) are very low.

[2] We also considered the alternative specification by reversing the direction of the regression in each case of (*a*), (*b*), and (*c*). This is known as normalization procedure. The cointegration test results obtained from this normalization consideration appear to be invariant to the specification presented in this paper and, therefore, are not reported.

Table 3
Results of the Bivariate Cointegrating Regressions

Between Price Variables:	Unit Root Test for the Residual			
	ADF Test With		Johansen LR Test	
	Constant	Constant and Trend	$r = 0$	$r = 1$
(*a*) $P_A = -0.163 + 0.84\ P_Y$; $R^2 = 0.44$ (13.80)	–3.31	–4.33	22.521	2.661
(*b*) $P_B = 1.078 + 0.59\ P_Y$; $R^2 = 0.37$ (11.89)	–7.38	–7.81	23.972	8.068
(*c*) $P_A = 0.074 + 0.58\ P_B$; $R^2 = 0.20$ (7.66)	–2.21	–3.83	23.956	2.697
Between Each Price and the Shock Variables:	**Unit Root Test for the Residual**			
	ADF Test With		Johansen LR Test	
	Constant	Constant and Trend	$r = 0$	$r = 1$
(*d*) $P_A = 1.30 - 0.55\ IP$; $R^2 = 0.005$ (–1.04)	–1.47	–3.99	39.229	8.124
(*e*) $P_B = 2.02 + 1.77\ IP$; $R^2 = 0.08$ (4.52)	–2.32	–3.77	61.642	4.251
(*f*) $P_Y = 1.69 + 0.86\ IP$; $R^2 = 0.02$ (2.06)	–2.92	–4.05	45.852	8.218

Note: T-ratios are in the parentheses. For critical values of ADF statistic see table 2. Critical values of LR test at 1% and 5% levels of significance and for $r = 0$ and $r = 1$ are 24.988 and 20.168, and 12.741, and 9.094, respectively (Johansen and Juselius 1990, p. 209).

As each price variable is cointegrated with the shock variable, we need to examine the strength of cointegrating relationships between the price variables and the shock variable using a multivariate framework to justify the degree of substitutability between the species.

Details of multivariate cointegrating regressions results are reported in table 4. For the first set of regressions (*a*), (*b*), and (*c*), the second type of ADF statistics are significant for regressions (*a*) and (*c*) and only marginally insignificant for regression (*b*) at the 5% level. On the other hand, LR statistics indicate the existence of two cointegrating vectors at the 1% level.

For the second set of regressions between price variables we find that only the second type of ADF test values are significant at the 5% level for regression (*d*). Both types of ADF test values are significant for regression (*e*) and insignificant for regression (*f*) at the conventional level. On the other hand, the LR test provides evidence of two cointegrating vectors at the 1% level for their combination and thereby leads us to reject the null hypothesis of non-cointegration.

Following the similar analysis for the third set of regressions, we find that either both, or at least one type of ADF test values are significant at the 5% level for all regressions (*g*), (*h*), and (*i*). The LR tests strongly support the existence of three cointegrating vectors in this case. Thus, the null hypothesis of non-cointegration can be rejected in favor of the alternative of cointegration.

To examine the strength of cointegrating relationships between variables, we compare the absolute value of the cointegrating parameters for each regression pre-

Table 4
Results of the Multivariate Cointegrating Regressions

Regressions:	Unit Root Test for Residual ADF Test With: Constant	Constant & Trend
(*a*) $P_A = -0.15 + 0.86\,P_Y - 1.29\,IP$; $R^2 = 0.47$ (14.41) (–3.32)	–1.92	–3.65
(*b*) $P_B = 1.06 + 0.57\,P_Y + 1.28\,IP$; $R^2 = 0.41$ (11.63) (4.06)	–2.83	–3.41
(*c*) $P_A = -0.01 + 0.65\,P_B - 1.70\,IP$; $R^2 = 0.24$ (8.52) (–3.54)	–1.13	–3.96

Johansen LR Test: Hypothesis	Regression (*a*)	(*b*)	(*c*)	Critical Values 1%	5%
H_O: $r = 0$	102.798	123.760	164.452	40.198	35.068
H_O: $r \leq 1$	31.247	33.802	45.742	24.988	20.168
H_O: $r \leq 2$	5.536	6.835	7.898	12.741	9.094

Regressions: (Between Price Variables)	Unit Root Test for Residual ADF Test With: Constant	Constant & Trend
(*d*) $P_A = -0.25 + 0.79\,P_Y + 0.08\,P_B$; $R^2 = 0.48$ (10.33) (0.97)	–1.95	–3.62
(*e*) $P_Y = 0.37 + 0.39\,P_A + 0.40\,P_B$; $R^2 = 0.57$ (10.33) (8.20)	–3.71	–3.69
(*f*) $P_B = 1.08 + 0.55\,P_Y + 0.05\,P_A$; $R^2 = 0.38$ (8.20) (0.97)	–2.56	–3.00

Johansen LR Test: Hypothesis	For Regressions (*d*), (*e*), and (*f*)	Critical Values 1%	5%
H_O: $r = 0$	153.839	40.198	35.068
H_O: $r \leq 1$	56.177	24.988	20.168
H_O: $r \leq 2$	5.073	12.741	9.094

Regressions: (All Prices and the Shock Variables)	Unit Root Test for Residual ADF Test With: Constant	Constant & Trend
(*g*) $P_A = -0.32 + 0.78\,P_Y + 0.15\,P_B - 1.49\,IP$; $R^2 = 0.48$ (10.39) (1.92) (–3.72)	–1.89	–3.65
(*h*) $P_Y = 0.39 + 0.40\,P_A + 0.39\,P_B + 0.41\,IP$; $R^2 = 0.57$ (10.39) (7.22) (1.39)	–3.68	–3.67
(*i*) $P_B = 1.08 + 0.48\,P_Y + 0.10\,P_A + 1.41\,IP$; $R^2 = 0.42$ (8.20) (1.92) (4.40)	–3.05	–3.44

Johansen LR Test: Hypothesis	For Regressions (*g*), (*h*), and (*i*)	Critical Values 1%	5%
H_O: $r = 0$	220.660	60.054	53.347
H_O: $r \leq 1$	101.492	40.198	35.068
H_O: $r \leq 2$	30.937	24.988	20.168
H_O: $r \leq 3$	5.498	12.741	9.094

Note: T-ratios are in the parentheses. For critical values of ADF statistics see table 2.

sented in tables 3 and 4. The strength of relationships between different combinations of the variables can be summarized by stating the range of estimated cointegrating parameters values obtained from different regressions as follows. For the combination (P_A, *IP*), (P_B, *IP*) and (P_Y, *IP*) the ranges are 0.55 to 1.49, 1.28 to 1.77, and 0.41 to 0.86 respectively. The ranges of parameter values for the combinations (P_A, P_B), (P_A, P_Y), and (P_B, P_Y) are 0.05 to 0.65, 0.39 to 0.86, and 0.39 to 0.59 respectively. It is noted that only for regression (*c*) in table 4 the cointegrating parameter for P_B is 0.65 which constitutes the highest value of the range for the combination (P_A, P_B). For other regressions in table 4 the parameter values for the same combination (P_A, P_B) are very low. The same is not true for other combinations.

To further evaluate our analysis, comparison should be made between the bivariate and the multivariate regressions results. Adding the shock variable to the first set of regressions in table 3, we find that R^2s do not increase substantially (highest 4%) and the change in the strength of co-movement (measured by relevant estimated parameters) between price variables is also marginal (see first set of regressions in table 4). On the other hand, adding one more price variable to the first set of regressions (*a*) and (*b*) in table 3, we find not only the change in R^2 values, but also the change in the strength of co-movement for combinations (P_A, P_Y) and (P_B, P_Y) are again negligible [see second set of regressions (*d*) and (*f*) in table 4]. Furthermore, for regression (*e*) in table 4 the strength of cointegrating relationships is almost the same as before for the combinations (P_A, P_Y) and (P_B, P_Y). Additionally, for that second set of regressions in table 4, we find that the strength of cointegrating relationships between the combinations (P_A, P_Y) and (P_B, P_Y) are substantially higher than the combination (P_A, P_B).

Comparing the second set of regressions with the third set in table 4, we find that the values of R^2 do not change at all for regressions (*g*) and (*h*) and a marginal (4%) increase for regression (*i*), and the strength of the relationships between the same combination of variables in second set of regressions (in table 4) follows the same patterns.

Thus, from the above results we find that: (*i*) although in most cases the coefficient estimates of the shock variable are statistically significant, overall, the variable does not have significant explanatory power in both bivariate and multivariate regressions, and (*ii*) the degree of substitutability between albacore and yellowfin, and bigeye and yellowfin is substantially higher than the degree of substitutability between albacore and bigeye. Moreover, our results from the multivariate regressions not only reflect the observable behavior of the consumers in the Japanese market but are also consistent with the data series. Therefore, it is reasonable to say that bigeye and yellowfin and, yellowfin and albacore could be appreciated as substitutes to each other in consumption. However, using yellowfin as a common link in these substitutability relationships, it is not correct to surmise that bigeye and albacore could also be treated as substitutes to each other. The reason is that consumers' perceptions, and hence preferences, regarding bigeye and albacore are affected by the ratings applied to the product's attribute categories. These are subsequently affected by socioeconomic and demographic factors (Engle and Kouka 1995; Nauman, *et al.* 1995). Upon examination of the price series it is noted that in most cases, price of bigeye is almost double the price of albacore. On the other hand, price differentials between bigeye and yellowfin and, yellowfin and albacore, are narrower (see figure 1). This larger price differential between bigeye and albacore reflects the difference in their attributes. Our argument is not only consistent with the observable market behavior but also in line with the published literature (see Wessells and Anderson 1992; Gordon, *et al.* 1993).

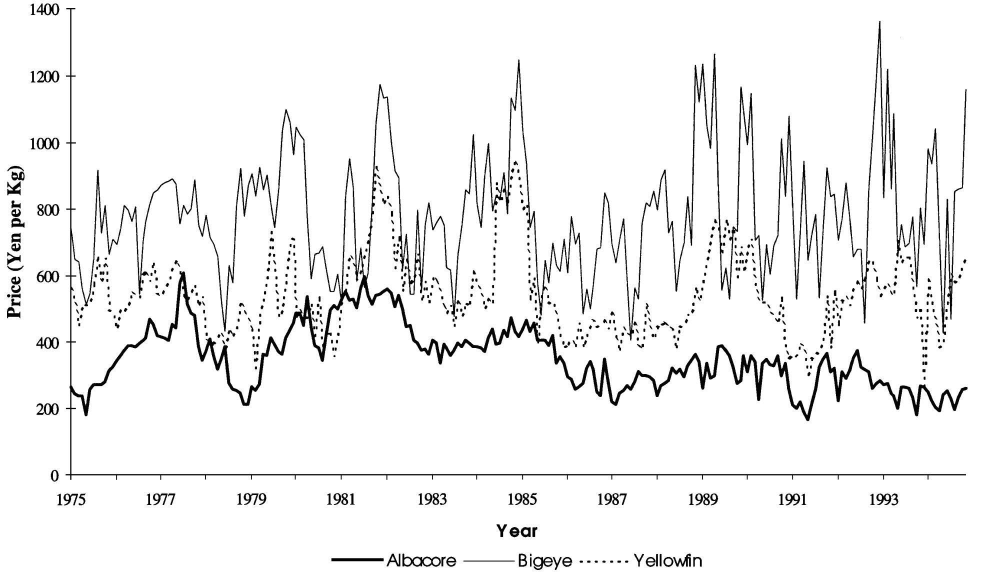

Figure 1. Monthly Average Nominal Prices of Bigeye, Yellowfin, and Albacore for the Period January, 1975 to November, 1994

Concluding Remarks

This paper applies recent techniques of cointegration to test statistical hypotheses concerning the existence, as well as the cause of long-run relationships between tuna species in the Japanese Yaizu fish market. A shock variable is introduced into the system to determine the degree of influence on the price relationships as well as the magnitude of the power in explaining the variation in prices of tuna species. Although, in most cases the coefficient estimates of the shock variable are statistically significant, but overall, the variable does not have significant explanatory power in both bivariate and multivariate regressions. We also find that the degree of substitutability between bigeye and albacore is substantially lower than the degree of substitutability between bigeye and yellowfin and, yellowfin and albacore. Thus, based on our empirical findings, we offer the following tentative conclusions: (*i*) bigeye and yellowfin and, yellowfin and albacore could be appreciated as substitute to each other, and (*ii*) bigeye and albacore can only be appreciated as substitute products when the end users' perception regarding the quality of the product is changed through proper handling, processing, promotional campaign, and pricing decision. We believe our findings provide a strong rationale for further study of price interrelationships, between different species, for the formulation of a sound marketing plan to meet the long-term economic objective of the seafood industry in Japan. It would also provide an understanding of the Japanese wholesale market price behavior which would be of great benefit to the resource-supplying nations of the South Pacific.

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